Differential Fertility, Human Capital, and Development

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Using micro-data from 48 developing countries, this article studies changes in cross-sectional patterns of fertility and child investment over the demographic transition. Before 1960, children from larger families obtained more education, in large part because they had richer and more educated parents. By century's end, these patterns had reversed. Consequently, fertility differentials by income and education historically raised the average education of the next generation, but they now reduce it. Relative to the level of average education, the positive effect of differential fertility in the past exceeded its negative effect in the present. While the reversal of differential fertility is unrelated to changes in GDP per capita, women's work, sectoral composition, or health, roughly half is attributable to rising aggregate education in the parents' generation. The data are consistent with a model in which fertility has a hump-shaped relationship with parental skill, due to a corner solution in which low-skill parents forgo investment in their children. As the returns to child investment rise, the peak of the relationship shifts to the left, reversing the associations under study.

Key words: Fertility, Human capital, Economic development, Economic growth

JEL Codes: O15, I25, J11, J13

1. INTRODUCTION

Over the last two centuries, most of the world's economies have seen unprecedented increases in living standards and decreases in fertility. Recent models of economic growth have advanced the understanding of the joint evolution of these economic and demographic processes.[1] Central to many, but not all, of these theories is the idea that a rising return to investment in children altered the calculus of childbearing, enabling the escape from the Malthusian trap. Although an abundance of aggregate time series evidence helps to motivate this work, efforts to understand the role of heterogeneity within an economy have been hampered by fragmentary evidence on how cross-sectional patterns of fertility and child investment change over the course of the demographic transition. Using a range of data covering half a century of birth cohorts from 48 developing countries, this article provides a unified view of how those patterns change, linking them to theories of the interplay between demography and economic growth.

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1. Collectively labeled "Unified Growth Theory" (Galor, 2011), these models have explored the roles of a variety of factors, including scale effects on technological progress (Galor and Weil, 2000), increases in longevity (Kalemli-Ozcan, 2002; Soares, 2005), changes in gender roles (Galor and Weil, 1996; Voigtländer and Voth, 2013), declines in child labour (Hazan and Berdugo, 2002; Doepke and Zilibotti, 2005), and natural selection (Galor and Moav, 2002).

Two strands in the theoretical literature relate to this focus on cross-sectional heterogeneity in fertility and skill investment during the process of growth.[2] The first, due to Galor and Moav (2002), analyses the evolutionary dynamics of lineages that have heterogeneous preferences over the quality and quantity of children.[3] A subsistence constraint causes fertility to initially be higher in richer, quality-preferring families, but as the standard of living rises above subsistence, fertility differentials flip. Consequently, in the early (Malthusian) regime, fertility heterogeneity promotes the growth of quality-preferring lineages, raising average human capital; in the late (modern) regime, it promotes the growth of quantity-preferring lineages, dampening human capital growth. A second strand in the literature—including papers by Dahan and Tsiddon (1998), Morand (1999), De la Croix and Doepke (2003), and Moav (2005)—fixes preferences and examines how the initial distribution of income or human capital interacts with fertility decisions to affect growth and income distribution dynamics.[4] These authors assume a specific structure of preferences and costs to reproduce two patterns observed in most present-day settings: (1) that wealthy parents have fewer children than poor parents and (2) that they educate their children more. As in Galor and Moav's (2002) modern regime, heterogeneity in fertility lowers average skill.[5] Indeed, much of this work posits that the higher fertility of the poor can help explain macroeconomic trends in developing countries during the post-war era. Interest in this idea dates back to Kuznets (1973), who conjectured that differential fertility adversely affects both the distribution and the growth rate of income.

But did rich or high-skill parents have low relative fertility in developing countries throughout this period? At least since Becker (1960), economists have recognized that although fertility decreases with income or skill in most settings today, the relationships may have once been positive. Along these lines, in the mid- to late-twentieth century, some small, cross-sectional studies in mostly rural parts of Africa and Asia showed a positive relation between fertility and parental income or skill (Schultz, 1981). Other studies of similar contexts revealed that children from larger families obtained more schooling, consistent with higher fertility among better-off parents (Buchmann and Hannum, 2001). Studies of historical Europe also suggest that fertility once increased with economic status in some settings.[6] But researchers have had to look far back in history to find these patterns in Western data, and economic historians are far from a consensus on whether they represent a regularity (Dribe *et al*., 2014). In the US, for example, the relationship has been negative for as long as measurement has been possible (Jones and Tertilt, 2008).

Efforts to form a unified view of these results have taken three approaches: (1) combining results from studies that use a variety of methods and measures (Cochrane, 1979; Skirbekk, 2008), (2) analysing survey data collected contemporaneously in several contexts (UN, 1987, 1995; Cleland and Rodríguez 1988; Mboup and Saha, 1998; Kremer and Chen, 2002), or (3) studying data from a single population over time (Maralani, 2008; Bengtsson and Dribe *et al*., 2014; Clark and Cummins, 2015). Although informative, these approaches are limited in their ability to clarify when the associations change; how those changes relate to theories of growth

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2. In an important contribution to this literature not directly related to growth, Mookherjee *et al*. (2012) derive conditions under which steady state reasoning disciplines the wage-fertility relationship to be negative.

3. See Clark (2007) and Galor and Michalopoulos (2012) for evolutionary theories emphasizing different sources of preference heterogeneity.

4. Althaus (1980) and Kremer and Chen (2002) consider similar issues in models that assume a specific relationship between parental skill and fertility, rather than allowing it to arise from parental optimization.

5. Several models also demonstrate how these fertility gaps can give rise to poverty traps, thus widening inequality. Empirically, Lam (1986) shows that the effect of differential fertility on inequality depends on the inequality metric, but his finding does not overturn the general equilibrium reasoning of recent theories.

6. See Weir (1995) and Hadeishi (2003) on France; Clark and Hamilton (2006) on Britain; and Bengtsson and Dribe *et al*. (2014) on Sweden.

and demographic transition; and what implications they have for the next-generation's human capital distribution.

This article seeks to fill that gap by analysing the evolution of two cross-sectional associations in many developing countries over many decades: (1) that between parental economic resources (proxied by durable goods ownership or father's education) and fertility and (2) that between sibship size and education. The results show that, in the not-too-distant past, richer or higher-skill parents had more children, and children with more siblings obtained more education. Today, the opposite is true for both relationships. These findings have implications for theories of fertility and the demographic transition, as well as for understanding the role of differential fertility in the process of growth. In particular, until recently, differences in fertility decisions across families raised the *per capita* stock of human capital instead of depressing it.

To guide the empirical work, the article begins by showing how skill differentials in fertility can change sign in the growth literature's standard framework for the study of cross-sectional fertility heterogeneity, due to De la Croix and Doepke (2003) and Moav (2005).[7] Within that framework, both papers assume that children cost time, while education costs money, which yields the negative gradient that is prevalent today. I demonstrate that with the addition of a subsistence constraint or a goods cost of children, the same framework predicts that fertility increases with income or skill among the poor, so that the relationship between parental resources and fertility is hump-shaped. In the early stages of development, when most parents have little income or skill, children with more siblings come from better-off families and obtain more education. With growth in income or skill, the association of parental resources with fertility turns negative. Additionally, in the model specification with a goods cost of children (but not with a subsistence constraint), a rising return to child investment moves the peak of the hump to the left, so that the association of parental resources with fertility can flip even without income growth.

The empirical analysis illustrates these results with two datasets constructed from the Demographic and Health Surveys (DHS). For the first, I treat survey respondents (who are women of childbearing age) as mothers, using fertility history data to construct two cross-sections of families from twenty countries in the 1986–94 and 2006–11 periods. In these data, respondents enumerate all of their children ever born, with information on survival status. Between the early and late periods, the associations of fertility with parental durable goods ownership and paternal education flipped from positive to negative in Africa and rural Asia; they were negative throughout in Latin America.[8] I argue that these patterns capture the tail end of a global transition from a positive to a negative association. Consistent with the existence of both goods and time costs of children, non-parametric estimates show that these relationships start hump-shaped, but the peak of the hump shifts to the left over time, eventually disappearing altogether, leaving a negative slope.

For the second dataset, I treat the DHS respondents as siblings, using sibling history data to retrospectively construct a longer panel of families from forty-three countries. In these data, respondents report all children ever born to their mothers, again with information on survival status. Among birth cohorts of the 1940s and 50s, most countries show positive associations between the number of ever-born or surviving siblings and educational attainment. Among cohorts of the 1980s, most countries show the opposite. The transition timing varies, with Latin America roughly in the 1960s, Asia roughly in the 1970s, and Africa roughly in the 1980s.

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7. Jones *et al.* (2010) discuss related theoretical issues but do not explore how these differentials reverse over time.

8. These results appear to contradict Kremer and Chen (2002) and their sources, which find in the same data that total fertility rates predominantly decrease with maternal education. Section 4.1 discusses reasons for the discrepancy and explains why this article's approach may be more fruitful for studying differential fertility over the long run.

Taken together, the data imply that in nearly all sample countries, the associations between parental economic resources and fertility and between sibship size and education both flipped from positive to negative. Although the DHS offers little data on childhood economic circumstance, three supplementary datasets (from Bangladesh, Indonesia, and Mexico) suggest that one can attribute much of the reversal in the sibsize–education association to the reversal of the link between paternal education and fertility.[9] In all, the analysis covers forty-eight countries, the richest being Mexico. Data from these forty-eight countries point to a broad regularity, though their implications for the historical development of more advanced economies are less certain.

To test alternative theories of this reversal, I assemble a country-by-birth cohort panel of sibsize–education coefficients. Net of country and cohort fixed effects, neither women's labour force participation, nor sectoral composition, nor GDP *per capita*, nor child mortality predicts the sibsize–education association. Rather, one variable can account for over half of the reversal of the sibsize–education association: the average educational attainment of the parents' generation. Because the reversal is uncorrelated with economic growth, its most likely cause is not a shift of the income distribution over the peak of a stable, hump-shaped income-fertility profile. Instead, much as the non-parametric fertility history results suggest, a rising return to investment in children may have lowered the income threshold at which families begin to invest, moving the peak of the income-fertility profile to the left. In many endogenous growth models, aggregate human capital raises the individual return to child investment, providing an explanation for the link between average adult education and differential fertility.

These findings imply changes in the effect of differential fertility on average human capital in the next generation, which I quantify relative to a counterfactual with an exogenous fertility level that applies to all families. The theoretical framework shows that one can separate this effect into two components, the first reflecting how equalizing fertility across families would affect the composition of the next generation, the second reflecting how it would affect families' child investment decisions. Model calibration reveals that the second component is extremely sensitive to the choice of the exogenous fertility level. I thus focus on the first component, which is invariant to the exogenous fertility level and is also estimable by means of a simple reweighting procedure. The procedure compares actual average educational attainment with the average that would arise if all families had the same number of children, with no change to their education (*i.e.* no reoptimization). In deriving this composition effect, the article contributes to a growing demographic literature on the aggregate consequences of cross-sectional associations (Mare and Maralani, 2006).

According to the results of the reweighting procedure, differential fertility raises average education in the early stages of development but decreases average education in the later stages. Since human capital investment is low in the early stages, the early positive effect of differential fertility is proportionally more important than the later negative effect. Indeed, among the least educated cohorts in the sample, differential fertility raises cohort mean education by as much as one-third and by 15% on average. If differential fertility plays an important role in growth, that role is thus most likely *positive*. The data also allow one to estimate, for women born around 1960, how differential fertility affected each country's distance to the world human capital frontier, which for that cohort is the US. In over 90% of sample countries, differential fertility reduced the shortfall in mean human capital relative to the US, with an average reduction of roughly 3% and a maximum reduction of roughly 9%.

By shedding light on the timing, causes, and consequences of the reversal of differential fertility in the developing world, this article contributes to several literatures. Most apparent is

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9. The supplementary datasets also indicate that the reversal of the sibsize–education association is similar for men (who are not in the DHS) and women.

the connection with two empirical literatures: one on parental socioeconomic status and fertility, the other on sibship size and education. In these literatures, evidence on positive associations is scattered, lacking a unifying framework.[10] This article uncovers a common time path in which both associations flip from positive to negative. Building on the theoretical growth literature, it provides a theoretical framework that explains the reversal and gives insight into its aggregate implications. Along these lines, the article shows how cross-family heterogeneity in fertility increases average education early in the development process but decreases it later; proportionally, the positive effect is much larger than the negative effect. That finding adds to our understanding of how demography interacts with the macroeconomy and calls attention to how cross-sectional patterns can inform models of fertility decline. The basic time-series facts about fertility decline are overdetermined, so a more thorough treatment of changing heterogeneity within populations will help narrow the field of candidate theories of the demographic transition.

2. CROSS-SECTIONAL PATTERNS IN A QUALITY–QUANTITY FRAMEWORK

This section studies how a subsistence constraint or a goods cost of children affect the growth literature's standard theoretical framework for studying differential fertility. Given the article's focus, I derive the model's cross-sectional properties and then briefly discuss its dynamic implications.

2.1. *Setup*

Parents maximize a log-linear utility function over their own consumption (c), the number of children (n), and human capital per child (h):

$$U(c,n,h) = \alpha \log(c) + (1-\alpha)(\log(n) + \beta \log(h)). \tag{2.1}$$

$\alpha \in (0,1)$ indexes the weight the parents place on their own consumption relative to the combined quantity and quality of children, while $\beta \in (0,1)$ reflects the importance of quality relative to quantity. Child quality, or human capital, is determined by:

$$h(e) = \theta_0 + \theta_1 e. \tag{2.2}$$

where e denotes education spending per child but also represents broader investment in children.[11] $\theta_0 > 0$ is a human capital endowment, while $\theta_1 > 0$ is the return to investment in children. The presence of θ_0 implies that the elasticity $\frac{eh'(e)}{h(e)}$ increases with e—which Jones et al. (2008) point out is crucial for an interior solution in which fertility declines with parental skill—and also allows for a corner solution with no child investment. θ_1 reflects both the return to skill and the price of skill, the latter being a function of school availability, teacher quality, and the opportunity cost of children's time.

  Irrespective of human capital, each child costs $\tau \in (0,1)$ units of time and $\kappa \geq 0$ goods. These costs represent the minimum activities (*e.g.* pregnancy, child care) and goods (*e.g.* food, clothing)

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10. A separate literature takes interest in the causal effect of exogenous increases in family size (Rosenzweig and Wolpin, 1980; Black *et al*., 2005; Li *et al*., 2008; Rosenzweig and Zhang, 2009; Angrist *et al*., 2010; Ponczek and Souza, 2012), where the evidence is also mixed.

11. The assumption that $\beta < 1$ plays no important conceptual role in the theory, but it guarantees the existence of a solution under a linear human capital production function. If one adds concavity to the production function, for example by setting $h(e) = (\theta_0 + \theta_1 e)^\sigma$ with $\sigma \in (0,1)$, then one can obtain a solution so long as β is smaller than $\frac{1}{\sigma} > 1$.

required for each child.[12] Parents are endowed with human capital H, drawn from a distribution $F(H)$ with support on $[\underline{H},\overline{H}]$, and w is the wage per unit of parental human capital. Thus, the budget constraint is:

$$c+\kappa n+ne\leq wH(1-\tau n). \tag{2.3}$$

The parents may also face a subsistence constraint, so that c must exceed $\tilde{c}\geq 0$.[13] The framework allows the goods cost of children and the subsistence level to be zero, in which case it becomes similar to the models of differential fertility by De la Croix and Doepke (2003) and Moav (2005). I seek to understand how its predictions change when either of these parameters is positive.

2.2. *Optimal fertility and child investment*

The framework yields closed-form solutions for optimal fertility and child investment. To characterize these solutions, two threshold levels of parental human capital are important. The first is $\widetilde{H}\equiv\frac{1}{\tau w}\left(\frac{\theta_0/\theta_1}{\beta}-\kappa\right)$, above which parents begin to invest in their children. If parental human capital is below $\widetilde{H}$, then parents are content with the human capital endowment θ_0, choosing a corner solution with no child investment. For higher skill parents, investment per child rises linearly in their human capital: $e_H^*=\frac{\beta(\kappa+\tau wH)-\theta_0/\theta_1}{1-\beta}$ if $H\geq\widetilde{H}$.

In addition to $\widetilde{H}$, fertility decisions depend on the threshold $\frac{\tilde{c}}{\alpha w}$, above which parents cease to be subsistence-constrained:

$$n_H^*=\begin{cases} \frac{wH-\tilde{c}}{\kappa+\tau wH} & \text{if } H<\min\left(\frac{\tilde{c}}{\alpha w},\widetilde{H}\right)\\ \frac{(1-\alpha)wH}{\kappa+\tau wH} & \text{if } \frac{\tilde{c}}{\alpha w}\leq H<\widetilde{H}\\ \frac{(1-\beta)(wH-\tilde{c})}{\kappa-\theta_0/\theta_1+\tau wH} & \text{if } \widetilde{H}\leq H<\frac{\tilde{c}}{\alpha w}\\ \frac{(1-\alpha)(1-\beta)wH}{\kappa-\theta_0/\theta_1+\tau wH} & \text{if } H\geq\max\left(\frac{\tilde{c}}{\alpha w},\widetilde{H}\right). \end{cases} \tag{2.4}$$
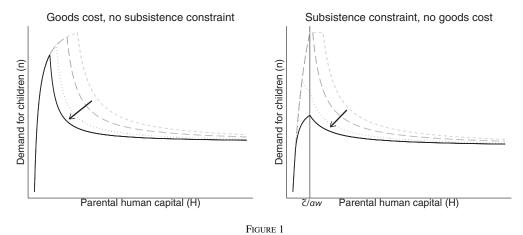
In the first line, parents are both subsistence constrained and at an investment corner solution. After consuming $\tilde{c}$, they spend all of their remaining full income wH on child quantity, so fertility increases with H. The next two lines deal with the cases in which $\frac{\tilde{c}}{\alpha w}<\widetilde{H}$ and $\widetilde{H}<\frac{\tilde{c}}{\alpha w}$, respectively. In the second line, the subsistence constraint no longer binds, but the parents remain at an investment corner solution. They devote αwH to their own consumption and the remainder to child quantity, so fertility is increasing in H if $\kappa>0$ and constant if $\kappa=0$. In the third line, the subsistence constraint binds, but the parents now choose an investment interior solution, making the comparative static ambiguous: $\frac{dn_H^*}{dH}\gtreqless 0$ if and only if $\kappa\gtreqless\frac{\theta_0}{\theta_1}-\tau\tilde{c}$. It is also ambiguous in the final line, in which the parents are constrained by neither the subsistence constraint nor the lower bound on child investment: $\frac{dn_H^*}{dH}\gtreqless 0$ if and only if $\kappa\gtreqless\frac{\theta_0}{\theta_1}$. If the goods cost is not too large, the substitution effect of a higher wage dominates the income effect.

To summarize, either a subsistence constraint or a goods cost of children guarantees a hump-shaped relationship between parental human capital and fertility, so long as the goods cost is not too large.[14] At low human capital levels, fertility increases with human capital if $\kappa>0$ or

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12. The model implicitly focuses on surviving children, abstracting from child mortality. One can view quantity costs τ and κ as incorporating the burden of mortality. I mainly address this issue empirically, showing in Section 5 that mortality decline is unrelated to the main results.

13. Assume that $\underline{H}>\tilde{c}$, so the lowest-skill parents can meet the subsistence constraint.

14. De la Croix (2013) and Murtin (2013) also discuss the role of a goods cost in generating a hump-shaped income-fertility profile.

FIGURE 1

Changes in the optimal fertility schedule as the return to child investment increases.

Notes: Assumes that $\kappa < \theta/\theta_1$, so the demand for children declines in parental human capital in the interior solution.

$\tilde{c} > 0$; at high human capital levels, it decreases with human capital if the goods cost is smaller than the ratio θ_0/θ_1. The same hump shape holds for income. Thus, this framework, based on homogenous preferences but heterogeneous initial skill, generates a skill-fertility profile similar to that in Galor and Moav's (2002) model, which combines preference heterogeneity with a subsistence constraint.

One can glean insight into the importance of goods costs vis-à-vis subsistence constraints by studying the response of the skill-fertility profile to an increase in the return to child investment. Rising skill returns are crucial to many economic models of the demographic transition, so this comparative static is key.[15] Figure 1 depicts how the relationship between parental human capital and fertility changes after successive increases in the return to child investment (θ_1). The two panels reveal how the framework's predictions depend on whether the hump shape is driven by a goods cost of children or a subsistence constraint. In the left panel, which assumes a positive goods cost of children but no subsistence constraint, increases in θ_1 shift the peak of the hump shape downward and to the left. Fertility falls among parents that are at an interior solution, and as $\widetilde{H}$ falls, more parents switch from a corner solution to an interior solution. In the right panel, which assumes a subsistence constraint but no goods cost, increases in θ_1 still depress unconstrained parents' fertility but have no systematic effect on the location of the peak.

Malthus (1826) posited that total output growth causes population growth, leading many authors to refer to a positive *aggregate* income-fertility link as "Malthusian" (Galor, 2011). One can reasonably extend this reasoning to view a positive *cross-sectional* association between parental resources and fertility as Malthusian. Two mechanisms are likely to shift a population from such a Malthusian regime to a modern fertility regime with a negative association. First, the distribution of full income (wH) could shift to the right, over the peak of the hump, because of rising wages or parental human capital. In this case, broad-based gains in living standards would tend to flip the association from positive to negative. Second, an increase in the return to child investment could shift the peak of the hump to the left, flipping the association even without rising incomes. This second mechanism is unambiguous only under a goods cost of children. The

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15. Taking a broad view of the skill returns, this statement applies to Becker *et al*. (1990), Dahan and Tsiddon (1998), Morand (1999), Galor and Weil (1996, 2000), Galor and Moav (2002), Doepke and Zilibotti (2005), Hazan and Berdugo (2002), Kalemli-Ozcan (2002), and Soares (2005), among others.

empirical work sheds light on these mechanisms by estimating the n_H^* profile non-parametrically and by analysing the aggregate determinants of changing socio-economic patterns in fertility.

2.3. *Aggregate implications*

To characterize the effect of differential fertility on average human capital, I consider a counterfactual in which fertility is exogenously fixed at $\tilde{n}$ children.[16] Under this level of exogenous fertility, parents with human capital H invest in their children as follows:

$$
e_H^{\tilde{n}} = \begin{cases} 0 & \text{if } H < \frac{\kappa\tilde{n}+\left(\frac{\alpha}{\beta-\alpha\beta}\right)\frac{\theta_0}{\theta_1}\tilde{n}}{w(1-\tau\tilde{n})} \\ \left(\frac{1}{\tilde{n}}-\tau\right)wH-\kappa-\frac{\tilde{c}}{\tilde{n}} & \text{if } \left(\frac{\alpha}{\beta-\alpha\beta}\right)\frac{\theta_0}{\theta_1}\tilde{n} \leq H < \frac{\left(1+\left(\frac{1-\alpha}{\alpha}\right)\beta\right)\tilde{c}+\kappa\tilde{n}-\frac{\theta_0}{\theta_1}\tilde{n}}{w(1-\tau\tilde{n})} \\ \frac{(\beta-\alpha\beta)\left(\left(\frac{1}{\tilde{n}}-\tau\right)wH-\kappa\right)-\alpha\theta_0/\theta_1}{\alpha+\beta-\alpha\beta} & \text{if } H \geq \max\left(\frac{\kappa\tilde{n}+\left(\frac{\alpha}{\beta-\alpha\beta}\right)\frac{\theta_0}{\theta_1}\tilde{n}}{w(1-\tau\tilde{n})}, \frac{\left(1+\left(\frac{1-\alpha}{\alpha}\right)\beta\right)\tilde{c}+\kappa\tilde{n}-\frac{\theta_0}{\theta_1}\tilde{n}}{w(1-\tau\tilde{n})}\right). \end{cases}
$$
(2.5)

The lowest-skill parents choose a corner solution with no child investment, whereas parents with intermediate skill invest but may be constrained by their subsistence requirements. For the highest-skill parents, only the budget constraint binds. Here as before, optimal child investment weakly increases in parental human capital. And consistent with a quality–quantity tradeoff, child investment decreases in $\tilde{n}$ when it is not at a corner solution.

The total effect of differential fertility is the difference in average human capital between endogenous and exogenous fertility:

$$
\Delta_{tot}(F,\tilde{n}) = \frac{\int h(e_H^*)n_H^*dF(H)}{\int n_H^*dF(H)} - \int h\left(e_H^{\tilde{n}}\right)dF(H).
$$
(2.6)

On the right-hand side of the equation, the first and second expressions equal average human capital under endogenous and exogenous fertility, respectively. To average across children rather than families, the first expression reweights the parental human capital distribution by the factor $\frac{n}{E[n]}$. The second does not because all families have the same fertility. The difference between these averages depends on the reweighting of the population *and* any change in investment behaviour.

In fact, one can decompose $\Delta_{tot}(F,\tilde{n})$ into quantities that reflect these two margins. To obtain this decomposition, add and subtract $\int h(e_H^*)dF(H)$, average human capital *across families*, to the right-hand side of Equation (2.6):

$$
\Delta_{tot}(F,\tilde{n}) = \underbrace{\int \left(\frac{n_H^*}{\int n_{\mathcal{H}}^*dF(\mathcal{H})}-1\right)h(e_H^*)dF(H)}_{\Delta_{comp}(F)} + \underbrace{\int \left\{h(e_H^*)-h\left(e_H^{\tilde{n}}\right)\right\}dF(H)}_{\Delta_{adj}(F,\tilde{n})},
$$
(2.7)

where $\mathcal{H}$ is a dummy of integration. $\Delta_{comp}(F)$ is the *composition effect* of differential fertility, which captures how fertility heterogeneity reweights the distribution of families from one generation to the next. Put differently, the composition effect measures how average human capital *across children* differs between the endogenous fertility optimum and the counterfactual

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16. Assume that $\tilde{n} < \frac{w\underline{H}-\tilde{c}}{\tau w\underline{H}+\kappa}$, so the exogenous fertility level does not keep the lowest-skill parents from meeting the subsistence constraint.

in which all families have an equal number of children but maintain the per child investments that were optimal under endogenous fertility. Because this counterfactual involves no re-optimization, the composition effect is invariant to $\tilde{n}$. $\Delta_{adj}(F,\tilde{n})$ is the *adjustment effect* of differential fertility, measuring how average human capital per child *across families* changes in response to the shift from endogenous to exogenous fertility. This component depends crucially on the exogenous fertility level, $\tilde{n}$. If $\tilde{n}$ were set to the smallest number of children observed in the endogenous fertility distribution, the adjustment effect would be positive; if it were instead set to the highest number of children, the adjustment effect would be negative. The empirical work focuses on the composition effect because it solely reflects the joint distribution of quantity and quality investments, rather than arbitrary exogenous fertility levels. The composition effect is also appealing because one can estimate it non-parametrically in any dataset with measures of family size and child investment.

Assuming a positive subsistence level and a small goods cost of children, several properties of the composition effect are apparent. If $\overline{H}<\widetilde{H}$, so that all parents make no educational investments, then $\Delta_{comp}(F)=0$. Growth in human capital, wages, or the return to child investment causes $\Delta_{comp}(F)$ to turn positive; fertility rates become highest in the small share of parents with positive child investment. As this process continues, more mass accumulates in the domain in which $\frac{dn_H^*}{dH}<0$, eventually turning $\Delta_{comp}(F)$ negative. Indeed, if $\underline{H}>\max\left(\frac{\tilde{c}}{\alpha w},\widetilde{H}\right)$, so that fertility decreases with parental human capital across the entire support of F, then $\Delta_{comp}(F)$ is unambiguously negative. These results suggest that in the early stages of economic development—when most are subsistence constrained or at an investment corner solution, but the wealthy few educate their children—the composition effect is positive. But with broad-based gains in living standards or increases in the return to child investment, the Malthusian fertility regime gives way to the modern fertility regime, and the composition effect turns negative.

These composition effects are generally relevant for economic growth, but growth effects may be especially pronounced with particular human capital externalities. For instance, if aggregate human capital raises the productivity of the education sector, then the composition effects can generate endogenous growth. This type of externality finds support in recent evidence that the quality of schooling is as important as the quantity of schooling in explaining cross-country variation in output per worker (Schoellman, 2012). If improvements in schooling quality raise θ_1 without affecting θ_0, then higher aggregate human capital in one generation causes greater educational investment in the next. In the Malthusian era, the positive composition effects of differential fertility raise the return to child investment, speeding both economic growth and the transition to negative composition effects (due to rising skills and leftward shifts in $\widetilde{H}$). Once negative composition effects set in, differential fertility retards growth (as in De la Croix and Doepke, 2003). This potential role for differential fertility in the emergence (and subsequent moderation) of modern growth is similar to the mechanism in Galor and Moav's (2002) model of evolution.

3. DATA ON TWO GENERATIONS OF SIBSHIPS

Using data from the DHS, I construct two generations of sibships by viewing respondents as mothers and daughters. Conducted in over ninety countries, the DHS interviews nationally-representative samples of women of childbearing age (usually 15–49 years).

3.1. *DHS fertility histories*

The first set of analyses draws on the fertility histories, in which respondents list all of their children ever born, with information on survival. I use these data to study how fertility relates

to paternal education and household durable goods ownership, a proxy for household wealth or income. Each of these measures has benefits and drawbacks. Paternal education is attractive because it measures parental human capital and is determined largely before fertility decisions. But its connection to fertility may go beyond the mechanisms in the theoretical framework, and its connection to full income changes with the wage rate.[17] Conversely, durable goods ownership provides a useful gauge of the household's economic resources, although it is an imperfect income proxy and may be endogenous to fertility decisions. And as with paternal education, relative price changes may complicate comparisons of durable goods ownership over time.

For a composite measure of durable goods ownership, I take the first principal component of a vector of ownership indicators for car, motorcycle, bicycle, refrigerator, television, and radio. This approach is similar to that of Filmer and Pritchett (2001), except that it does not incorporate measures of housing conditions (*e.g.* access to piped water), which may be communally determined. I perform the principal components analysis on the whole sample, so the resulting measure (which is standardized to have mean 0 and standard deviation 1) reflects the same quantity of durable goods in all countries and time periods.

Three sample restrictions are worthy of note. First, to avoid the complicated task of disentangling cohort effects from changes in the timing of childbearing, I focus on women at least 45 years old and interpret their numbers of children as completed fertility. The focus on older women also has the advantage of capturing cohorts of mothers more likely to be in the early regime in which fertility is increasing in income and skill. Second, because I analyse paternal education, I include only ever-married women (who report their husbands' education).[18] Third, I compare results from two time periods, pre-1995 and post-2005, and only include countries with survey data (including the relevant variables) from both periods, leaving 58,680 ever-married women from forty-six surveys in twenty countries. Appendix Table 1 lists countries and survey years.

3.2. *DHS sibling histories*

In some surveys, the DHS administers a sibling history module to collect data on adult mortality in settings with poor vital registration. The module asks respondents to list all children ever born to their mothers, with information on sex, year of birth, and year of death if no longer alive. In addition to adult mortality, the sibling histories offer a window into the sibling structure that adult women experienced as children. I relate this information to their educational attainment.

Most DHS surveys with sibling histories are representative of all women of childbearing age, but a few (from Bangladesh, Indonesia, Jordan, and Nepal) include only ever-married women. From these surveys, I minimize concerns about selection by only including age groups in which the rate of ever marriage is at least 95%. As a result, I include women over 30 years from the relevant surveys in Bangladesh and Nepal, but I discard such surveys from Indonesia and Jordan, where marriage rates are lower.[19] The analysis sample comprises eighty-three surveys from forty-three countries. To exclude respondents who have not finished schooling or whose mothers have

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17. The choice of paternal education is meant to strengthen the link to the theory, not to diminish the role of maternal education. Even more than paternal education, maternal education may affect preferences, beliefs, and bargaining power, and, because of low rates of female labour force participation, its link with income and the opportunity cost of time is tenuous. Section 4.1 ends with a description of results for maternal education.

18. The durable goods results are similar for all women and for ever-married women.

19. The relevant surveys are the 2001 Bangladesh DHS; 1996 Nepal DHS; 1994, 1997, 2002, and 2007 Indonesia DHS; and 1997 Jordan DHS. Later surveys in Nepal and Indonesia included never-married women and are thus included in the analysis. I also discard data from the 1989 Bolivia DHS and the 1999 Nigeria DHS due to data irregularities.

not completed childbearing, I drop data on women less than 20 years old, leaving 845,594 women born between 1945 and 1989. Appendix Table 1 lists countries and survey years.

3.3. *Supplementary surveys*

The DHS data are useful in their breadth but suffer from four shortcomings. First, they omit men, for whom the relationship between sibship size and education may be different. Second, they offer little information on the respondent's childhood environment, such as her parents' education. To supplement the DHS on these first two fronts, I draw on three surveys: the Indonesia Family Life Survey, the Matlab Health and Socioeconomic Survey, and the Mexico Family Life Survey. Third, DHS data do not include parental wages, complicating a direct mapping from the data to the theory. To fill this gap, I use an additional Indonesian dataset, the 1976 Indonesia Intercensal Population Survey. Fourth, the DHS do not include industrialized countries, so I use data from the US National Longitudinal Survey of Youth for comparison.

4. CHANGING CROSS-SECTIONAL FERTILITY PATTERNS

This section investigates the evolution of differential fertility in developing countries since the 1940s. It begins with fertility history data, analysing the socioeconomic determinants of fertility, and then turns to sibling history data, analysing the association of sibship size with completed education. All analyses use sampling weights, but the results are similar without them.[20]

4.1. *Fertility patterns by durable goods ownership and paternal education*

To assess the changing links between parental economic resources and fertility, I begin with a series of non-parametric estimations. Figures 2–3 show local linear regressions of completed fertility on measures of household economic resources, with accompanying kernel density estimates for the measures of economic resources. These estimations work together to illuminate the theory. The local linear regression estimates allow one to assess whether the resource–fertility relationship changes shape, while the kernel density estimates reveal whether the distribution of resources shifts underneath the relationship. Either phenomenon may flip the overall association of economic resources with fertility. Because individual surveys offer limited sample sizes of women over the age of 45 years, the figures pool countries in the same world region, on the premise that they have common economic fundamentals. Regions are ordered by increasing average paternal education.

Figure 2 reveals pronounced variation in the relationship between durable goods ownership and fertility, both across regions and within regions over time. In the least educated region, West Africa, both ever-born and surviving fertility monotonically increase in durable goods ownership during the early period, but the relationships flatten and even become slightly negatively-sloped in the late period. In most other regions, the relationships are initially hump-shaped but then become everywhere negatively-sloped. One can interpret this disappearance of the hump-shape as a shift in its peak towards resource levels lower than the sample minimum, so these results are consistent with a rising return to investment in children. The relationship's changing shape can alone flip the association of durable goods ownership with fertility, but such a flip would be bolstered by rising durable goods ownership in all regions, demonstrated by the kernel density

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20. In the fertility history analyses, which pool multiple countries, the sampling weights are re-scaled to sum to national populations in 1990 (for the pre-1995 sample) and 2010 (for the post-2005 sample).
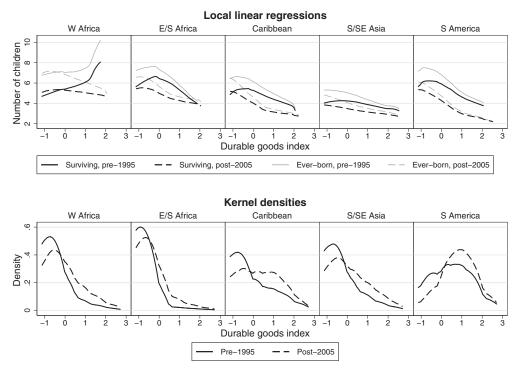
Local linear regressions

FIGURE 2

Durable goods ownership and completed fertility, ever-married women 45–49 years.

Notes: Regions are ordered by increasing average paternal education. Bandwidth is set to 0.5. Sample includes ever-married women in countries with a full durable goods module in both the early and late periods. Local linear regression sample (but not kernel density sample) omits the top 1% of the region/period-specific durable goods index distribution. The durable goods index is the first principal component of a vector of ownership indicators for car, motorcycle, bicycle, refrigerator, television, and radio. Data source: DHS Fertility Histories

estimates. That is to say, even if the relationship had remained hump-shaped, changes in the distribution of durable goods ownership would have alone made the association more negative. Nevertheless, changes in functional form would likely be the main driver of a reversal in the association of economic resources and fertility.

Similar patterns appear in Figure 3, which plots the relationship between paternal education and fertility. In the less-educated regions (Africa and the Caribbean), the relationship is initially hump-shaped but then turns negative; in Asia and South America, it is always negative. As with the distribution of durable goods ownership, the distribution of husband's education shifts to the right in all regions between the early and late periods. So here again, changes in both functional form and the distribution of the independent variable may promote a flip in the association of paternal education and fertility.

Thus, consistent with a rising return to child investment, all regions exhibit downward and leftward shifts in the peaks of the relationships between economic resources and fertility. A concern for this interpretation is that the relative prices of parental skill and of durable goods changed between the early and late periods. But the relative prices of these variables probably moved in opposite directions. On the one hand, cheaper consumer durables would imply that parents with a given level of durables ownership are poorer, making them more likely to be on the increasing segment of the hump shape. On the other hand, increased returns to skill
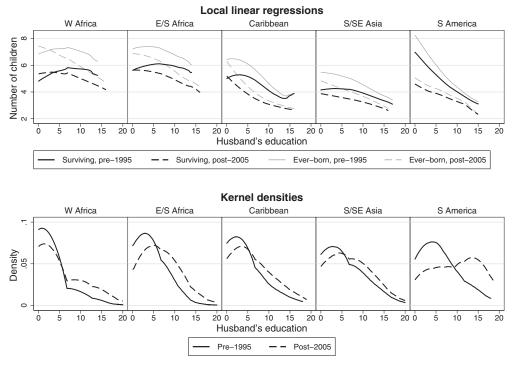
Figure 3

Paternal education and completed fertility, ever-married women 45–49 years.

Notes: Regions are ordered by increasing average paternal education. Bandwidth is set to 3. Sample includes ever-married women in countries with a full durable goods module in both the early and late periods. Local linear regression sample (but not kernel density sample) omits the top 1% of the region/period-specific paternal education distribution. Data source: DHS Fertility Histories

would tend to move parents of a given skill level towards the declining segment of the hump shape. Both variables point to similar patterns, mitigating concerns about the confounding role of prices.

In both figures, changing functional forms are evident for both ever-born fertility and surviving fertility. Which of these measures provides a better representation of the demand for children depends on the extent to which parents target surviving fertility. Given that fertility at ages 45–49 years reflects sequential childbearing decisions and deaths over three decades, it seems reasonable to interpret surviving fertility as a closer proxy for the demand for children.[21] Moreover, only surviving fertility is relevant for the composition effects estimated in Section 6; children who do not survive to adulthood do not affect the next generation's skill distribution. For conciseness, the remaining analyses focus on counts of surviving children only.

To quantify these changes in functional form and assess their statistical significance, Tables 1–2 fit ordinary least squares (OLS) regressions of surviving fertility on the measures of parental economic resources. The top panel of each table reports a linear specification to summarize whether, on average, better-off parents exhibit higher or lower fertility. To address

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21. Empirical research finds that parents "replace" deceased children rapidly (Ben-Porath, 1976), while theoretical work suggests that single-period models of the demand for children have similar quantitative implications to sequential models with child mortality (Doepke, 2005).

TABLE 1

Durable goods ownership and surviving fertility, ever-married women 45–49 years

	W. Africa		E./S. Africa		Caribbean		S./S.E. Asia		S. America	
	(Burkina Faso, Cameroon, Ghana, Niger, Nigeria, Senegal)		(Burundi, Kenya, Madagascar, Malawi, Namibia, Tanzania, Zambia, Zimbabwe)		(Dominican Republic, Haiti)		(India, Indonesia)		(Colombia, Peru)	
	'86–'94 (1)	'06–'11 (2)	'86–'94 (3)	'06–'11 (4)	'86–'94 (5)	'06–'11 (6)	'86–'94 (7)	'06–'11 (8)	'86–'94 (9)	'06–'11 (10)
A. Linear Specification										
Durable goods index	0.43 [0.15]*	−0.124 [0.046]*	0.004 [0.102]	−0.60 [0.07]*	−0.34 [0.10]*	−0.86 [0.06]*	−0.10 [0.03]*	−0.31 [0.02]*	−0.86 [0.10]*	−0.95 [0.04]*
p-value for difference	0.0004		<0.0001		<0.0001		<0.0001		0.40	
B. Quadratic Specification										
Durable goods index	0.46 [0.16]*	−0.042 [0.054]	0.01 [0.09]	−0.54 [0.07]*	−0.20 [0.11]	−1.01 [0.077]*	−0.037 [0.034]	−0.32 [0.030]*	−0.79 [0.15]*	−1.35 [0.068]*
Squared index	−0.23 [0.11]*	−0.14 [0.04]*	−0.72 [0.09]*	−0.17 [0.06]*	−0.32 [0.10]*	0.23 [0.050]*	−0.18 [0.026]*	0.019 [0.018]	−0.087 [0.089]	0.28 [0.031]*
Arg max of quadratic	1.01 [0.50]*	−0.15 [0.23]	0.01 [0.06]	−1.60 [0.65]	−0.30 [0.23]	—	−0.10 [0.10]	—	−4.53 [5.29]	—
p-value for difference	0.035		0.013							
Mean index	−0.6	−0.1	−0.8	−0.5	−0.3	0.2	−0.4	0.0	0.4	0.8
Mean num. of kids	5.0	5.2	5.9	5.3	5.0	4.0	4.1	3.5	5.3	3.3
Observations	2,772	8,247	2,777	4,885	1,132	3,112	11,682	13,715	2,356	8,002

Note: Brackets contain standard errors clustered at the primary sampling unit level. Sample includes ever-married women in countries with a full durable goods module in both the early and late periods. The durable goods index is the first principal component of a vector of ownership indicators for car, motorcycle, bicycle, refrigerator, television, and radio. The arg max is the ratio of the linear term to the quadratic term; its standard error is computed using the delta method. The symbol "—" indicates a convex quadratic function, which has no maximum. Data source: DHS Fertility Histories. *coefficient significant at the 5% level.

TABLE 2
Husband's education and surviving fertility, ever-married women 45–49 years

	W. Africa (Burkina Faso, Cameroon, Ghana, Niger, Nigeria, Senegal)		E./S. Africa (Burundi, Kenya, Madagascar, Malawi, Namibia, Tanzania, Zambia, Zimbabwe)		Caribbean (Dominican Republic, Haiti)		S./S.E. Asia (India, Indonesia)		S. America (Colombia, Peru)	
	'86–'94 (1)	'06–'11 (2)	'86–'94 (3)	'06–'11 (4)	'86–'94 (5)	'06–'11 (6)	'86–'94 (7)	'06–'11 (8)	'86–'94 (9)	'06–'11 (10)
A. Linear Specification										
Husband's education	0.087 [0.026]*	−0.057 [0.007]*	0.036 [0.019]	−0.09 [0.01]*	−0.10 [0.02]*	−0.17 [0.01]*	−0.034 [0.006]*	−0.068 [0.005]*	−0.26 [0.02]*	−0.13 [0.01]*
p-value for difference	<0.0001		<0.0001		0.001		<0.0001		<0.0001	
B. Quadratic Specification										
Husband's education	0.29 [0.08]*	0.08 [0.02]*	0.19 [0.04]*	0.01 [0.03]	0.120 [0.063]	−0.34 [0.04]*	0.05 [0.02]*	−0.05 [0.01]*	−0.40 [0.06]*	−0.10 [0.025]*
Squared education	−0.018 [0.006]*	−0.010 [0.002]*	−0.015 [0.003]*	−0.008 [0.002]*	−0.017 [0.004]*	0.012 [0.002]*	−0.006 [0.001]*	−0.0014 [0.0008]	0.010 [0.004]*	−0.0016 [0.0013]
Arg max of quadratic	7.94 [0.67]	3.93 [0.59]	6.08 [0.60]	0.65 [1.86]	3.53 [1.09]	—	3.81 [0.64]	−17.1 [14.2]	—	−32.6 [35.7]
p-value for difference	<0.0001		0.005				0.14			
Mean index	−0.6	−0.1	−0.8	−0.5	−0.3	0.2	−0.4	0.0	0.4	0.8
Mean num. of kids	5.0	5.2	5.9	5.3	5.0	4.0	4.1	3.5	5.3	3.3
Observations	2,772	8,247	2,777	4,885	1,132	3,112	11,682	13,715	2,356	8,002

Note: Brackets contain standard errors clustered at the primary sampling unit level. Sample includes ever-married women in countries with a full durable goods module in both the early and late periods. The arg max is the ratio of the linear term to the quadratic term; its standard error is computed using the delta method. The symbol "—" indicates a convex quadratic function, which has no maximum. Data source: DHS Fertility Histories. *coefficient significant at the 5% level.

the theory of the hump-shaped relationship, the bottom panel reports quadratic models. If the coefficient on the squared term is negative—consistent with a hump shape—then the bottom panel also provides the arg max of the quadratic function, estimated as the ratio of the linear term to the quadratic term.[22] These estimates allow one to test whether the peak of the hump-shape has shifted over time.

In the top panels of the tables, the linear association between parental resources and fertility becomes significantly more negative in four out of the five regions, irrespective of the measure of parental resources. The changes are starkest in Africa, where the association is initially positive. In West Africa, it flips from significantly positive to significantly negative, while in Eastern and Southern Africa, it shifts from insignificantly positive to significantly negative. The other three regions exhibit negative associations in both periods. Supplementary Appendix Table 2 separates the sample into urban and rural areas, finding more positive coefficients in rural areas; the coefficients go from positive to negative in rural parts of Western Africa, Eastern/Southern Africa, the Caribbean, and South/Southeast Asia. These results suggest a shared process that visits urban areas before rural, first in Latin America, then in Asia, and finally in Africa.

In the bottom panels of the tables, the quadratic specifications reveal significantly negative squared term coefficients—indicating hump-shaped relationships—in the early period in four out of the five regions. In all four of these regions (and for both measures of economic resources), the arg max of the estimated function lands within the support of the distribution of economic resources. By the late period, however, the functional form changes dramatically. Either the arg max shifts significantly to the left or the relationship loses its hump-shape altogether, implying an arg max that is infinitely negative. In some cases, the coefficient on the squared term even turns significantly positive, which is consistent with the asymptotic properties of the skill-fertility profile in the theory. Here again, Latin America appears to have already arrived at the modern fertility regime in the early period.

These results may appear to contradict much existing research, which finds in the same data that total fertility rates decrease with maternal education in most settings.[23] The difference in results has three likely sources. First, these other studies analyse total fertility rates, which mix the fertility behaviour of older and younger cohorts; I focus on the completed fertility of the older cohorts, thus capturing an earlier phase of the fertility transition. Second, the other sources only look at all children ever born, while I consider counts of ever-born children and surviving children. The changing patterns are observable for both measures of fertility, but they are clearer for counts of surviving children because child mortality is higher among the poor. Third, the other sources focus on maternal education, whereas I examine paternal education and durable goods ownership, which bear a stronger link with income and the opportunity cost of time, due to low rates of female labour force participation in many settings. For comparison with the existing literature, Supplementary Appendix Table 3 includes the durables index, paternal education, and maternal education in the same regression. Conditional on durable goods ownership and paternal education, maternal education is negatively associated with fertility in both periods.

4.2. Sibship size and educational attainment

The fertility history results provide evidence that the relationship between parental economic status and fertility was once hump-shaped, with a peak that shifted downward and leftward over

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22. The standard error of the arg max estimate is computed using the delta method.
23. See UN (1987, 1995), Cleland and Rodríguez, (1988) Mboup and Saha (1998), and Kremer and Chen (2002).

time. But they do not give a sense of how pervasive these changes were historically—whether they occurred in Latin America, for example. Furthermore, they do not offer any information over the implications of changing fertility differentials for average human capital. The sibling histories offer a window onto the issues for birth cohorts going back to the 1940s. Unfortunately, the DHS collects little data on economic conditions in childhood, but we can gain some insight into the evolution of the socioeconomic fertility differentials by studying changes in the relationship between sibship size and education. The sibsize–education link is also directly relevant for assessing the effect of differential fertility on the skill distribution.

To capture the long-run evolution of this association, I estimate regressions separately by country and 5-year birth cohort (1945–1949 to 1985–1989).[24] For woman i born in country c and cohort t, the regression specification is:

$$highest\,grade_{ict} = \delta_{ct} + \gamma_{ct} sibsize_{ict} + \varepsilon_{ict}, \tag{4.1}$$

where $highest\,grade_{ict}$ denotes her schooling and $sibsize_{ict}$ denotes her sibship size.[25]

Focusing only on surviving sibship size for tractability, Figure 4 displays estimates of γ_{ct} over time within each country. Positive sibsize–education associations were pervasive until recently but have now largely disappeared. Specifically, for earlier birth cohorts, most coefficients are significantly positive, while for the latest birth cohorts, few coefficients are significantly positive, and many are significantly negative. Consistent with the fertility history results, this reversal in the sibsize–education relationship occurs earliest in Latin America, followed soon thereafter by several countries in Asia. Africa's reversal is quite recent; several countries remain in the pre-reversal regime. The Andean countries provide the starkest examples. For Bolivian women born in the late 1940s, each additional sibling is associated with 0.4 *more* years of education, whereas for those born in the 1980s, an additional sibling is associated with 0.5 *fewer* years of education. For Peruvian women, the associations swing from 0.2 to −0.6 over the same period.

Figure 4 reports findings for surviving sibship size because it bears a closer link to the composition effect and, arguably, the demand for children. But results for ever-born sibship size are similar. Supplementary Appendix Figure 1 plots the association between siblings ever born and completed education, again revealing coefficients that start broadly positive but become negative over time. The within-country, between-country, and overall correlations between the ever-born coefficients and the surviving coefficients are all at least 0.95. Nevertheless, a possible concern for the interpretation of these results is that an adult respondent may not remember all of her deceased siblings, some of whom may have died before she was born.

4.3. *Connecting the results*

Results from the supplementary surveys in Bangladesh, Indonesia, and Mexico complement these findings in two ways. First, these surveys can address questions about gender heterogeneity, which the DHS cannot because it gathers sibling history data only from women. Since many countries exhibit different patterns of investment in boys and girls, the exclusion of men leaves open the question of whether the association of family size with investment in boys changes in similar ways to the association of family size with investment in girls. The supplementary surveys interview

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24. For precision, I omit cells with fewer than 200 observations, representing 2% of all cells.

25. Researchers often control for birth order in estimating the effect of family size on educational attainment. However, the present article is concerned not with causal effects but with equilibrium differences between large and small families, making regression adjustment unnecessary. Regardless, in unreported analyses, controlling for birth order did not substantively affect the results.
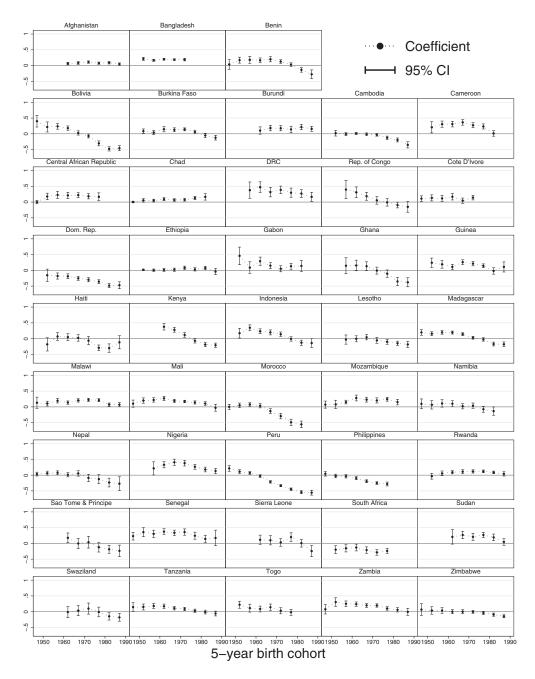
FIGURE 4

Association of surviving sibship size with education by period of birth.

Notes: From regressions of years of education on surviving sibship size. Sample includes all women over the age of 20 years, except for the 2001 Bangladesh survey and the 1996 Nepal survey, which include ever-married women over the age of 30 years. Data source: DHS Sibling Histories

both genders, allowing an investigation of this issue. In Supplementary Appendix Table 4, all three supplementary surveys show declining sibsize–education relationships for both genders. Changes in the association of family size with child investment are not specific to girls.

Second, the supplementary surveys can help connect the fertility history results with the sibling history results. The fertility history results seem to contain the last phases of the global transition to a negative relationship between parental economic status and fertility, while the sibling history results point to a widespread shift of the sibsize–education link from positive to negative. While the two phenomena seem connected, the absence of childhood background characteristics for adult respondents in the DHS prevents examination of this issue. The supplementary surveys, however, include data on paternal education. Mirroring the DHS fertility history results, Figure 5 shows a hump-shaped relationship between paternal education and surviving sibship size in all three countries, with the peak shifting to the left over time. This shift even occurs in the middle-income context of Mexico, which had higher GDP *per capita* than all of the DHS countries throughout the sample period, suggesting that the DHS results are broadly representative of a phenomenon that spanned the developing world.[26] Meanwhile, with extremely few exceptions, educational attainment monotonically rises in father's education. In other words, higher-skill parents always educated their children more, but over time, they shifted from high to low relative fertility. Along these lines, Supplementary Appendix Table 5 shows that the evolution of the sibsize–education relationship has much to do with the changing relationship between paternal education and sibship size. Within each country, sibsize–education coefficients decrease across successive birth cohorts, but controlling for paternal education mutes the reductions by half.

5. MECHANISMS OF THE REVERSAL

Viewed through the lens of the theoretical framework, the results so far are most consistent with an increase in the return to investment in children. The data point to flips in the overall associations of economic status with fertility and of sibship size with education, both from positive to negative. Estimated non-parametrically, the relationship linking parental economic resources with fertility is hump-shaped, with the peak moving downward and to the left over time. These patterns match the predictions of the theoretical framework under a goods cost of children and a rising return to investment in children. However, the reversal of differential fertility occurred during a half-century of much economic and demographic change, offering a further testing ground for this hypothesis. This section further explores the mechanisms of the reversal, first by analysing the aggregate determinants of differential fertility and then by discussing how it relates to leading alternative theories of demographic change.

5.1. *Aggregate determinants of the reversal*

This section estimates how changes in economic and demographic aggregates relate to changes in the sibsize–education association, γ_{ct}, which I treat as a summary statistic for differential fertility. I focus on the sibsize–education association rather than the resource–fertility association because the former offers a longer time horizon and is more precisely estimated at the country level. As in Section 4.2, I report results only for surviving sibship size, but unreported results for ever-born sibship size coefficients are qualitatively similar.

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26. The Mexico data also have counts of siblings ever born, which in unreported results also exhibit an initial hump shape.
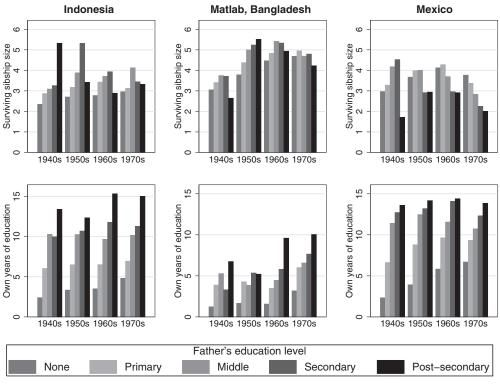
FIGURE 5

Average sibship size and education by decade of birth and father's education level, supplementary surveys.

Notes: Means are weighted by the survey weight divided by the surviving sibship size to make them representative of the families of surviving children. Data source: men and women born 1940–1979 in the Indonesia Family Life Survey (1993, 1997 waves), Matlab Health and Socioeconomic Survey (1996), and Mexico Family Life Survey (2002 wave)

5.1.1. Hypotheses. The rising-returns hypothesis makes several predictions regarding the aggregate correlates of γ_{ct}. First and second, a rise in the return to investment in children weakly decreases fertility and increases educational investment for all families. As a result, decreases in γ_{ct} should be associated with declining average family size and rising average child investment. Third, in the presence of intergenerational human capital externalities, decreases in γ_{ct} should be associated with rising average education among adults. As discussed in Section 2.3, such externalities would arise if more educated populations have more productive teachers.

The literature on the demographic transition cites several other forces as potentially important in driving fertility change. One such force is income growth, which may reverse fertility gaps by pushing poor parents over the peak of the hump. This force fails to explain the changing form of the hump, but it also offers the testable implication that decreases in γ_{ct} should be associated with growth in living standards. A second force is the improving circumstance of women. For example, rising female labour force participation may flip fertility differentials by increasing the correlation between household income and the opportunity cost of childbearing. Alternatively, expansions in female education may alter bargaining power or norms in a way that reverses fertility differentials. One can test these hypotheses with data on female labour force participation and female educational attainment. Two final forces that arise often in discussions of demographic change are mortality decline and urbanization, which one can examine directly in the data.

5.1.2.   Method.   I merge the panel of coefficients from Figure 4 with data on economic and demographic aggregates in the period of birth. Using the estimates of γ_{ct} as the dependent variable, I estimate:

$$\hat{\gamma}_{ct} = Z'_{ct}\lambda + \tau_t + \mu_c + \varepsilon_{ct}, \tag{5.1}$$

where Z_{ct} is a vector of economic and demographic aggregates, and τ_t and μ_c are cohort and country fixed effects, respectively. Residuals ε_{ct} are clustered at the country level. The specification nets out global trends and time-invariant country characteristics, thus allowing one to assess which aggregate changes are associated with the reversal of differential fertility. As a benchmark, one can measure the average magnitude of this reversal by leaving Z_{ct} out of equation (5.1). The resulting cohort effect estimates are flat through the early 1960s, at which point they begin a downward trend, becoming significantly negative in the 1970s. Net of country fixed effects, the sibsize–education association is on average 0.26 lower for the 1985–9 cohort than for the 1945–9 cohort.

5.1.3.   Results.   The first set of estimations, appearing in Table 3, defines the covariates Z_{ct} as cohort average outcomes from the DHS: average completed education, average surviving sibship size, and the average fraction of siblings dying before they reach the age of 5 years. Because these average outcomes are co-determined with the sibsize–education relationship, one should think of the estimands as equilibrium associations rather than causal effects. For this reason, I include only one covariate in each regression (in addition to the cohort and country fixed effects). Also, because the estimates of γ_{ct} and the cohort average outcomes are based on the same data, the table supplements the OLS results with estimations that correct for correlated measurement errors using Fuller's (1987) method-of-moments technique. To ensure that the standard errors fully capture variability from the first- and second-stage estimations, I bootstrap the entire procedure, sampling *within* individual surveys before running the first stage and then sampling *across* countries before running the second stage. The OLS results and the Fuller results give the same three conclusions: (1) as the sibsize–education association declines, average educational investment increases; (2) as the sibsize–education association declines, average family size declines; and (3) the sibsize–education association has no relation to child mortality rates. These findings support the rising-returns hypothesis while casting doubt on any role for mortality decline or health improvement.

Table 4 regresses $\hat{\gamma}_{ct}$ on three socioeconomic aggregates in the period of birth: log GDP *per capita*, average adult educational attainment, and urbanization. Data on these variables come from a variety of sources. GDP *per capita* is from from the Penn World Table (Heston *et al.*, 2012); average adult (ages 25+ years) educational attainment is from Barro and Lee (2010) and Cohen and Soto (2007); and urbanization is from UNDP (2011). To provide a summary measure of aggregate conditions at the time of fertility and child investment decisions, I average these variables over the 5-year period of birth.[27] Since the education measure comes from two datasets that do not completely overlap, the table presents one regression for a combined sample and a separate regression for each of the source samples. In the combined sample, I use the Barro–Lee estimates when available, and for countries that only have Cohen–Soto estimates, I use them to

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27. GDP *per capita* is available annually, but the Barro–Lee education data and the UNDP urbanization data are available in 5-year increments, while the Cohen–Soto education data are available in 10-year increments. For these sources, I linearly interpolate across years before taking 5-year averages. Unreported results based on the closest measurement rather than interpolation are similar.

TABLE 3

Demographic correlates of the sibship size-education relationship

	Mean (SD) (1)	OLS (2)	Fuller (3)	OLS (4)	Fuller (5)	OLS (6)	Fuller (7)
Cohort average education	4.3 (2.8)	−0.045 [0.019]*	−0.050 [0.023]*				
Cohort average surviving sibship size	4.4 (0.7)			0.100 [0.042]*	0.106 [0.052]*		
Cohort average fraction of siblings dying under 5 years	0.10 (0.04)					0.44 [0.79]	0.67 [2.21]
Number of observations	323	323	323	323	323	323	323
Number of countries	43	43	43	43	43	43	43

Note: All analyses include birth cohort and country fixed effects. Fuller estimates correct for measurement error from sampling variability. The dependent variable is the coefficient from a regression of education on surviving sibship size. Brackets contain bootstrapped standard errors. The bootstrap procedure resamples respondents from each survey before estimating country-cohort parameters and then resamples entire countries before estimating the analyses reported in the table Data source: DHS Sibling Histories. *coefficient significant at the 5% level.

TABLE 4

Development and the sibship size–education relationship

	OLS (1)	OLS (2)	OLS (3)
Ln(GDP per capita in birth period)	0.030 [0.074]	0.014 [0.076]	0.065 [0.069]
Avg. adult yrs. ed. in birth period	−0.101 [0.027]*	−0.113 [0.031]*	−0.128 [0.028]*
Fraction urban in birth period	−0.33 [0.45]	−0.27 [0.47]	−0.10 [0.36]
Number of observations	227	203	162
Number of countries	39	35	30
Education dataset	Combined	Barro–Lee	Cohen–Soto

Note: All regressions include birth cohort and country fixed effects. The dependent variable is the coefficient from a regression of education on surviving sibship size. Brackets contain standard errors clustered at the country level. The "combined" measure uses the Barro–Lee data if available and otherwise projects the Cohen–Soto data onto the Barro–Lee scale. *coefficient significant at the 5% level.

generate predicted Barro–Lee estimates based on a regression of Barro–Lee on Cohen–Soto. All three regressions lead to the same conclusion: while aggregate income growth and urbanization do not play a role, the rising educational attainment of the parent generation is intimately connected with the reversal of the sibsize–education relationship among offspring. In fact, the coefficient of −0.09 on average education implies that rising education can account for roughly 60% of the of 1985–9 cohort effect for γ_{ct}, as reported at the start of this section. Again, these findings best match the rising-returns hypothesis, while discrediting theories centered on income growth and urbanization.

Several alternative theories deal with the position of women; these theories are the subject of Table 5. Data on the variables in Table 5 are only available for small subsamples, so to maximize sample size, each regression uses a different sample. The table first addresses theories involving the expansion of women's labour market opportunities. Goldin (1995) argues that women's work becomes incompatible with child rearing only when it moves outside the home, such that agricultural work is less likely to compete with parenting than service work. If women's opportunity cost of time explains the reversal, then the emergence of the service sector must

TABLE 5
Female empowerment and the sibship size–education relationship

	OLS (1)	OLS (2)	OLS (3)
Women's labour force participation rate in birth period	0.045 [0.101]		
Manufacturing fraction of value added in birth period		0.06 [0.21]	
Services fraction of value added in birth period		−0.02 [0.23]	
Avg. adult male yrs. ed. in birth period			−0.071 [0.021]*
Avg. adult female yrs. ed. in birth period			−0.051 [0.033]
p-value: joint test of education coefficients			0.002
p-value: difference of education coefficients			0.851
Number of observations	115	144	246
Number of countries	35	41	35

Note: All regressions include birth cohort and country fixed effects. The dependent variable is the coefficient from a regression of education on surviving sibship size. Brackets contain standard errors clustered at the country level. The education measures in column (3) are from the Barro–Lee dataset. *coefficient significant at the 5% level.

also play a key role. To explore the role of the opportunity cost of women's time, Columns (1) and (2) thus regress $\hat{\gamma}_{ct}$ on the women's (ages 20–59 years) labour force participation rate (from ILO 2012 and Olivetti, 2014) and the sectoral composition of value added (from Heston *et al*., 2012).[28] Neither variable plays a role in the reversal of the sibsize–education association. Another gender-specific theory emphasizes female education over male. Column (3) thus uses gender-disaggregated education data, available only from the Barro–Lee dataset, to ask whether the role of average education is due to women or men. While the coefficients on average female education and average male education are jointly significantly different from zero, they are not significantly different from each other; in fact, the coefficient on average male education is larger and individually more significant. Table 5 suggests that the causes of the reversal are not specific to the empowerment of women.

5.2. *Relation to other theories of demographic change*

Combined with the non-monotonic patterns in Section 4, the aggregate panel results favour the hypothesis that rising returns to investment in children reversed fertility differentials by income and skill. Income growth, urbanization, mortality decline, and women's empowerment appear to have played little role in the reversal. However, the literature on the demographic transition—especially in disciplines outside economics—suggests some alternative theories of the reversal that the results so far do not directly address.

In interpreting the changing cross-sectional patterns, many non-economists would think first of preferences, rather than prices or returns. Several theories of fertility decline (Caldwell, 1980, 1982; Casterline, 2001) posit changes in beliefs and norms regarding child-rearing, and some versions of these theories could explain the observed regime change. Consider the introduction of new "Western" norms that increase the relative importance of child quality in the utility function

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28. Sectoral composition is available on an annual basis, but female labour force participation is measured sporadically, so I linearly interpolate country values between measurements. Unreported analyses using only un-interpolated values led to the same conclusions.

(β), raising optimal education and lowering optimal fertility. If these norms affect the richest (or most educated) families most strongly, then the relationship between fertility and income (or skill) could flip from positive to negative, starting at the right tail of the income distribution. Caldwell, 1980, 1982) assigns much importance to mass education in altering childbearing norms, thus predicting a relationship between γ_{ct} and average adult educational attainment. Without further structure, the theory is otherwise difficult to test. But if one views the norms as increasing the social return to education spending (θ_1), then the norms theory coincides with the rising-returns hypothesis.

An alternative version of the norms theory associates the diffusion of new norms with the empowerment of women. If women have higher β's than men, and if women of higher income or education make the earliest gains in household bargaining power, then richer households will be the first to transition to low fertility. Unlike the broader norms theory, one can reject the role of female empowerment on the basis of the results in Table 5. Neither women's work nor women's schooling plays an important role in the reversal of differential fertility.

A separate class of theories emphasizes upward intergenerational transfers from children to parents, in the form of child labour or old-age support.[29] In the theoretical framework of Section 2, the falling prevalence of child labour has similar consequences to a rising return to child investment. Some of the decline in child labour may actually be the result of increases in the return to child investment. Some might also be due to new sanctions against child labour, which one could characterize as increases in the goods cost of children κ. Just as with an increase in the return to child investment, an increase in the goods cost of children decreases the wage threshold at which families start to spend on education, which can shift the peak of the fertility hump to the left. This mechanism is complementary to the return-to-investment hypothesis.

Other theories stressing intergenerational transfers do not fit as well into Section 2's framework. Under an old-age security motive for childbearing, for example, financial deepening could flip the income–fertility relationship if wealthy families substituted other savings vehicles for children. But this reasoning gives no account for why decreases in quantity investment would be accompanied by increases in quality investment. Furthermore, as stressed by Galor (2011), wealthier couples typically have access to a wider variety of savings vehicles before the fertility transition. Finally, Lee (2000) argues no society exhibits a net upward flow of resources across generations, unless one counts the pension systems of rich countries. Based on these insights, the old-age support motive is unlikely to explain the reversal of differential fertility.

Nevertheless, some specific institutional arrangements do give rise to significant pecuniary returns to having many adult offspring. For example, Tertilt (2005) posits that men in polygynous societies have an incentive to invest their wealth in a large number of children. In such societies, a groom typically "buys" a bride from her father, so men benefit from having many daughters but do not lose from having many sons. Consistent with this idea, the average sibsize–education association in polygynous countries exceeds that in monogamous countries by 0.1 to 0.2, both within Africa and across the world.[30] But declining polygyny cannot explain the global patterns documented in Section 4 because both polygyny and bride price were rare in many sample countries throughout the study period.

Instead of the prices or returns associated with childbearing, advocates of family planning might emphasize the uneven adoption of effective contraceptive technology (Potts, 1997). From this perspective, the currently negative relationship between income and fertility is due to an unmet need for contraception among the poor. But a theory of this type fails to account for the

29. See Cain (1983), Nugent (1985), Ehrlich and Lui (1991), Morand (1999), and Boldrin and Jones (2002).

30. Note that the patterns here must be driven by the number of children per wife, not the number of wives per husband. The DHS sibling history asks for siblings with the same biological mother.

early regime during which fertility increases in income. One possibility is that women from richer households have a higher biological capacity to bear children due to their better health. But this theory receives little support from the data, which show little relation between health improvements (as measured by child mortality rates) and the reversal of differential fertility. Theories that involve changes in the demand for children appear to provide a better fit to the data.

6. DIFFERENTIAL FERTILITY AND AVERAGE EDUCATION

The results so far suggest that differential fertility once promoted aggregate human capital accumulation rather than hindering it. This section delves into the magnitudes of these effects. Recall that equation (2.7) separated the effect of differential fertility on average human capital into two components: a composition effect, reflecting how fertility heterogeneity reshapes the population in the next generation, and an adjustment effect, reflecting how the change from exogenous to endogenous fertility alters parental optimization. To give a quantitative sense of both composition and adjustment effects on child investment, it first calibrates the model from Section 2 using an Indonesian dataset from the 1970s. It then returns to the DHS data, presenting non-parametric estimates of the composition effect on average education.
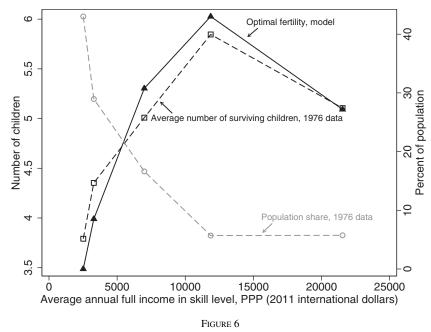
6.1. *Model calibration of composition and adjustment effects*

Because the datasets used in Sections 4–5 do not offer information on parental wages, I draw on yet another, the 1976 Indonesia Intercensal Population Survey (SUPAS), for model calibration. SUPAS is one of few developing-country datasets of its era with a large sample and information on labour income and hours worked. I calibrate the model to match patterns of surviving fertility among married women aged 45–54 year with husbands in wage- or salary-paying jobs. Because many of the respondents to the Indonesia Family Life Survey were born to this cohort of women, one can view these fertility patterns as mapping roughly onto the supplementary survey results in Figure 5. I separate husbands into five skill categories (none, incomplete primary, complete primary, complete middle, complete secondary and above) and define full annual income as the average hourly wage in each skill category times 16 hours per day times 365 days. Full annual income rises with skill and averages \$5,104, measured in PPP-adjusted 2011 international dollars.

To parametrize the model, I first set the time cost τ, followed by the preference parameters α and β, the goods cost κ, and finally the parameters of the human capital production function θ_0 and θ_1. I assume no subsistence constraint. In choosing τ, I follow De la Croix and Doepke (2003), who set it at 0.075 based on evidence from Haveman and Wolfe (1995) and Knowles (1999) that a child costs her parents 7.5% of their adult time endowment. I calibrate the preference parameter α to minimize the distance between each skill group's optimal fertility in the model and its average number of surviving children in the data, with equal weight given to each skill group. Given the parameter choices below, this approach leads to $\alpha = 0.438$, implying that parents spend 44% of their full income on their consumption. One can then back out β with an assumption regarding the limit of optimal fertility n_H^* as H approaches infinity. I assume that asymptotic fertility $n_\infty^* = \frac{(1-\alpha)(1-\beta)}{\tau}$ equals 2, which implies $\beta = 0.733$.[31]

One can additionally use α to set an appropriate goods cost of children. Using data from Indonesia during the same period, Muellbauer and Deaton (1986) apply equivalence scale methods

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31. The assumption that fertility approaches 2 as H approaches infinity may appear to be at odds with current fertility patterns in Europe and East Asia, where total fertility rates are lower than 2. However, for the relatively poor populations under consideration, the assumption that fertility approaches replacement levels seems reasonable.

FIGURE 6

Model calibration, Indonesia.

Notes: Sample includes 1246 married women aged 45–54 years with husbands in wage- or salary-paying jobs. Each point represents a skill group based on the husband's education: from left to right, none, incomplete primary, complete primary, complete middle, and complete secondary and above. Data source: SUPAS 1976

to conclude that a child costs 30–40% of an adult. I choose the midpoint of this range and follow De la Croix and Doepke (2003) in supposing that parents live with their children for half of their adult lives. Assuming that the equivalence scale estimate holds for parents at the median full income level, and that parents at this full income level do not invest in their children, I thus set κ to be 0.175 times α times median full income, or \$217 per year of adulthood. This value lands between the child costs implied by Indonesia's urban and rural per capita poverty lines for 1976.

Finally, the values of θ_0 and θ_1 only affect the solution through the ratio $\frac{\theta_0}{\theta_1}$, which one can obtain by combining the preceding parameter choices with an assumption regarding the peak of the n_H^* curve. Among the five skill groups in the data, the highest average number of surviving children is 5.8, but because the n_H^* curve reaches its maximum between two skill levels, I assume a peak fertility level of 6.5 to add flexibility to the model.[32] This assumption implies that $\frac{\theta_0}{\theta_1} = 1198$.

This parametrization predicts an average fertility rate of 4.78, relative to 4.82 in the data. Figure 6 displays each skill group's population share and average fertility in the data, as well as the model's predictions. As in Figure 5's results for Indonesia, fertility is lowest among the lowest-skill parents, who comprise the largest share of the population. It is highest at intermediate skill levels. Relative to the data, the model slightly understates fertility at low skill levels, slightly overstates fertility at intermediate skill levels, and matches fertility among the highest-skilled parents. Starting from this baseline parametrization, Table 6 presents a series of computational experiments.

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32. Assuming lower peak fertility does not alter model's predictions regarding differential fertility and its effects, but it leads the model to understate mean fertility.

TABLE 6
Computational experiments, Indonesia

	Average fertility ($\bar{n}$)	Education spending		Effect of differential fertility on education spending per child [Expressed as a share of column (2)]				
		per child	as share of tot. spending	Composition effect	Adjustment effects			
					$\tilde{n} = n_{min}$	$\tilde{n} = \bar{n}$	$\tilde{n} = n_{max}$	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	
Baseline	4.78	$38	3%	18%	−216%	−28%	48%	
Effect of 10% rise in…								
θ_1 only	4.63	$57	5%	5%	−128%	0%	67%	
w only	4.78	$59	5%	4%	−110%	4%	68%	
Both θ_1 and w	4.67	$74	6%	−9%	−76%	20%	73%	
Effect of 20% rise in…								
θ_1 only	4.54	$70	6%	−6%	−97%	15%	73%	
w only	4.80	$75	5%	−9%	−64%	22%	82%	
Both θ_1 and w	4.66	$95	7%	−31%	−105%	43%	73%	

Note: Parameter choices described in text. Education spending per child is denominated in PPP-adjusted 2011 international dollars and reflects average annual spending over parents' adult lives. Sample includes 1246 married women aged 45–54 years with husbands in wage- or salary-paying jobs. Data source: SUPAS 1976.

The first row of Table 6 further summarizes the baseline parametrization. As shown in columns (2)–(3), education spending amounts to $38 per child, or 3% of total expenditure. The remaining columns of the table report the effect of differential fertility on education spending per child, expressed as a share of the quantity in column (2). I focus on education spending rather than human capital because it is the closest analogue to educational attainment, which we observe in the data. Column (4) reports the composition effect, which raises education spending per child by 18% of the $38 reported in column (2).

In columns (5)–(7), the focus shifts to adjustment effects for different values of exogenous fertility $\tilde{n}$, which vary widely. Compared to a counterfactual in which fertility is exogenously set at the lowest observed fertility level, endogenous fertility reduces parents' average educational investments per child by over 200% of the $38 reported in column (2). Since the lowest-skill parents exhibit the lowest fertility, the low exogenous fertility level forces all other parents to have fewer children, so that educational investment is substantially higher in the counterfactual. In contrast, when the counterfactual sets fertility at its maximum observed level, the adjustment effect becomes positive, at 48%. When fertility is exogenously set at its average in the counterfactual, the effect shrinks to −28%.

The remaining rows of Table 6 consider how 10–20% increases in investment returns and wages affect average fertility, average education spending, and the effects of differential fertility. As predicted in Section 2, an increase in θ_1 alone reduces average fertility, raises average education spending, and reduces the composition effect of differential fertility, which eventually turns negative. While an increase in w alone has similar effects on average education spending and the composition effect, it also raises average fertility because most families are on the increasing segment of the fertility profile. Because the aggregate panel results suggest that fertility *decline* typically accompanies the flip of the association between fertility and child investment, the data better fit a rise in θ_1. However, the calibration results also show that simultaneous increases in θ_1 and w also reduce average fertility, raise average education spending, and reduce the composition effect. Throughout all of these computational experiments, the adjustment effect continues to vary wildly with the exogenous fertility level.

Although the article has focused on a single, exogenous fertility level as its main thought experiment, other counterfactuals are also of interest. For example, De la Croix and Doepke (2003) argue that when fertility monotonically declines with skill, higher inequality reduces education spending per child.[33] But this conclusion can change dramatically with a hump-shaped relationship between fertility and skill. Holding fixed the mean of the full income distribution, a 10% increase in its standard deviation raises education spending per child by roughly one-half, to \$58, in the baseline parametrization. Conversely, a mean-preserving, 10% decrease in spread reduces education spending per child by more than one-half, to \$16.

Although changes in variance are informative, some may be more interested in the effects of implementable policies. A 10% tax on labour income that is redistributed in the form of a lump-sum transfer reduces education spending per child by even more, to \$8. Compared to the mean-preserving reduction in spread, this redistributive policy has stronger negative effects because it reduces the opportunity cost of time and raises non-labour income, both of which raise fertility and decrease child investment. If the tax revenue is returned in the form of a child tax credit, the investment disincentive becomes so strong that no parents invest in their children. Another policy instrument of interest is a one-child policy, which dramatically limits parents' choice sets, reducing fertility by 60–83%. Such a policy would bind far more than China's famous law, which affected only a portion of the population and was implemented after fertility had already declined by half over the preceding decade, to three children per woman (Hesketh *et al*., 2005). In any case, this sharp contraction in fertility leads to a 41-fold increase in education spending per child, primarily reflecting re-optimization rather than compositional changes. Importantly, these quantities and the adjustment effects in Table 6 ignore any general equlibrium effects.

6.2. *Non-parametric estimates of the composition effect*

Composition effects in the calibrated model begin at 18% of education spending per child and decline to mildly negative values with moderate increases in θ_1 or w. Because the composition effect is a function of the realized joint distribution of fertility and child investment, one can verify these magnitudes in the data. Equation (2.7) expressed the composition effect by integrating over the parental wage distribution. I only observe siblings, with little information about their parents, so the formula for the composition effect is not directly estimable. Applying the law of iterated expectations, however, one can rewrite it over the distribution of surviving sibship sizes:

$$\Delta_{comp}(F) = \sum_{k=1}^{K} \left(\eta_k - \frac{\eta_k/k}{\sum_{l=1}^{K} \eta_l/l} \right) \mu_k, \tag{6.1}$$

where K is the maximum possible sibship size, η_k is the share of the individuals from surviving sibships of size k, and μ_k is the mean human capital (or educational investment) of individuals from sibships of size k. Inside the parentheses, the term η_k weights the sample to give mean human capital across individuals, while the term $\frac{\eta_k/k}{\sum_{l=1}^{K} \eta_l/l}$ reweights the sample to give mean human capital (or educational investment) across families. Attractively, this expression captures *any* composition effect of heterogeneity in fertility and skill investment, not just the heterogeneity

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33. De la Croix and Doepke find that higher inequality reduces the next generation's *human capital*, which results from both the concavity of their human capital production function and changes in education spending per child.

specific to the model in Section 2. Nevertheless, the reweighted mean omits childless couples. This omission is not an issue in the context of the model, in which no couples choose to be childless. But in practice, the counterfactual average represents a world in which the offspring of actual parents, rather than potential parents, are equally represented in the next generation. I use the empirical analogues of η_k and μ_k to estimate $\hat{\Delta}_{comp}$ and obtain its variance using the delta method.

I begin with estimates of the composition effect on average educational investment, which bears the closest link to the calibration results in Section 6.1. For successive 5-year birth cohorts within each country, Figure 7 displays estimates of the composition effect of differential fertility on mean educational attainment. The results overturn the conventional wisdom that variation in fertility over the skill or income distribution tends to lower mean education. Only South Africa exhibits a significantly negative composition effect for the earliest cohort. In some countries, predominantly African, differential fertility increased mean educational attainment throughout the sample period. Most of the remaining countries have undergone a transition from a regime in which differential fertility promoted the growth of mean education to a regime in which differential fertility depressed it. For two compelling examples, consider the Andean nations of Bolivia and Peru. In the 1945–9 cohort, differential fertility increased mean education by 0.3 to 0.5 years in both countries; in the 1985–9 cohort, differential fertility reduced average education by 0.5 years.

One can shed light on the importance of these magnitudes by expressing each cohort's composition effect as a share of its mean educational attainment. Figure 8 carries out this exercise, with the composition effect as a share of mean education on the vertical axis and mean education on the horizontal axis. A local linear regression (with a bandwidth of 1 year) captures how the relative importance of the composition effect changes with rising mean education. At low levels of mean education, the composition effect is proportionally large and positive, raising mean education by as much as one-third and by 15% on average. This result bears striking similarity to the baseline calibration result in Indonesia. As mean education increases, the composition effect trends downward, becoming negative when mean education reaches the secondary level. Proportionally, these negative composition effects are small, averaging 5% of mean education. Contrary to the assumptions of much of the literature on differential fertility and economic growth, composition effects are *positive* when they are proportionally most important. The 15% positive effect may or may not be large enough to play an important role in endogenous growth, but it is more likely to do so than the 5% negative effect.

Another way to place these composition effects in context is to calculate the extent to which differential fertility affects cross-country skill gaps. Data are available for all sample countries for the birth cohorts of the 1960s and early 1970s. According to Barro and Lee (2010), the US leads the world in average female education for these cohorts. To study the effects of differential fertility at the frontier of female human capital, I thus use data from the National Longitudinal Survey of Youth 1979 (NLSY79), which covers US birth cohorts from 1958 to 1965. Among women in the NLSY79, the composition effect is −0.17 years of education, amounting to 1.3% of mean educational attainment. The NLSY79 thus suggests that the composition effect remains proportionally small even at mean education levels higher than those observed in DHS cohorts.

Table 7 combines the NLSY79 and the DHS to ask how differential fertility affects each DHS country's distance to the human capital frontier for women of the 1958–65 birth cohort. For reference, columns (1) and (2) report actual and reweighted mean educational attainment; the difference between the two is the composition effect. In this early birth cohort, the composition effect is positive for most countries. Columns (3)–(5) then estimate how much larger or smaller is a country's distance to the human capital frontier as a result of differential fertility, compared to
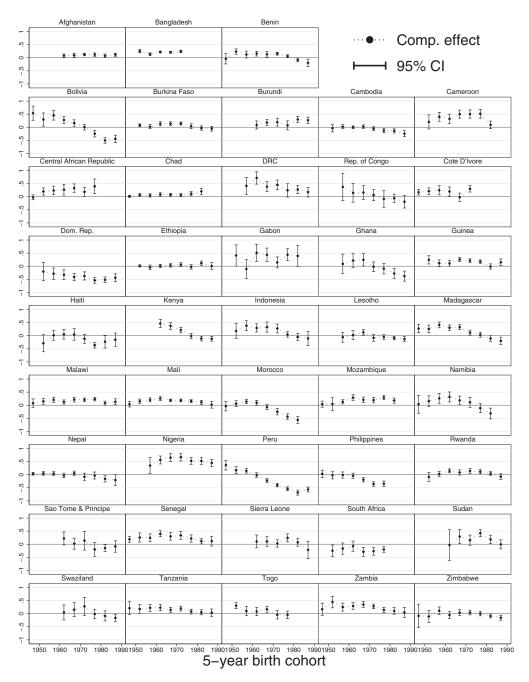
FIGURE 7

Composition effects of differential fertility by period of birth.

Notes: Difference between average education and the counterfactual in which all families had the same number of siblings, with no change to their education. CIs calculated with the delta method. Sample includes all women over the age of 20 years, except for the 2001 Bangladesh survey and the 1996 Nepal survey, which include ever-married women the age of 30 years. Data source: DHS Sibling Histories

TABLE 7
Differential fertility and distance to the human capital frontier, women born 1958–1965

| | Rewt. | Comp. effect of diff. fert. on distance to frontier in… | | |
| | | | | |
	Avg. ed. (1)	avg. ed. (2)	Avg. ed. (3)	Avg. HC (4)	Agg. HC (5)
Afghanistan	0.43	0.36	−1.8%	−2.3%	−2.1%
Nepal	0.66	0.65	−1.4%	−1.9%	−1.9%
Burkina	0.67	0.54	−2.2%	−2.7%	−2.5%
Ethiopia	0.67	0.67	−1.2%	−1.8%	−1.8%
Chad	0.72	0.65	−1.8%	−2.3%	−2.2%
Mali	1.04	0.80	−3.1%	−3.5%	−3.2%
Burundi	1.05	1.02	−1.5%	−2.1%	−2.0%
Guinea	1.09	0.94	−2.4%	−2.9%	−2.7%
Benin	1.48	1.36	−2.3%	−2.8%	−2.7%
Sudan	1.56	1.03	−5.4%	−5.8%	−5.0%
Mozambique	1.65	1.53	−2.3%	−2.8%	−2.7%
Senegal	1.82	1.44	−4.4%	−4.7%	−4.3%
Sierra Leone	1.91	1.68	−3.2%	−3.7%	−3.4%
Central African Rep.	2.08	1.82	−3.6%	−3.9%	−3.7%
Cote D'Ivore	2.10	1.97	−2.5%	−3.0%	−2.9%
Bangladesh	2.15	1.97	−3.0%	−3.4%	−3.3%
Togo	2.25	2.24	−1.6%	−2.1%	−2.1%
Haiti	2.43	2.30	−2.6%	−3.0%	−3.0%
Cambodia	2.44	2.43	−1.5%	−2.1%	−2.1%
Morocco	2.44	2.35	−2.3%	−2.8%	−2.7%
Rwanda	2.62	2.59	−1.7%	−2.3%	−2.3%
Malawi	2.88	2.73	−2.9%	−3.4%	−3.4%
Nigeria	3.95	3.65	−4.6%	−5.1%	−5.3%
DRC	4.04	3.69	−5.0%	−5.4%	−5.7%
Madagascar	4.04	3.70	−5.0%	−5.3%	−5.7%
S. Tome & Principe	4.06	3.89	−3.4%	−3.8%	−4.0%
Tanzania	4.12	3.83	−4.6%	−4.9%	−5.1%
Cameroon	4.39	3.99	−5.7%	−6.0%	−5.7%
Ghana	5.18	4.86	−5.4%	−5.7%	−5.6%
Zimbabwe	5.61	5.61	−2.1%	−2.6%	−2.8%
Zambia	5.62	5.30	−5.8%	−6.1%	−6.1%
Kenya	6.04	5.61	−7.3%	−7.5%	−7.7%
Indonesia	6.23	5.88	−6.5%	−7.0%	−7.1%
Bolivia	6.29	6.17	−3.8%	−4.2%	−4.5%
Namibia	6.34	6.30	−2.7%	−3.3%	−3.4%
Swaziland	6.39	6.58	0.2%	−0.8%	−0.6%
Gabon	6.66	6.27	−7.3%	−7.9%	−8.1%
Rep. Of Congo	6.66	6.17	−8.5%	−8.8%	−9.3%
Lesotho	6.77	6.83	−1.6%	−2.5%	−2.4%
Dominican Rep.	7.61	8.04	4.4%	3.7%	4.3%
South Africa	7.94	8.19	1.4%	0.4%	−0.1%
Peru	8.05	8.11	−1.9%	−2.5%	−3.2%
Philippines	8.90	8.94	−2.5%	−3.3%	−3.8%
Average	**3.74**	**3.60**	**−3.1%**	**−3.6%**	**−3.6%**

Note: Countries ordered by average education. The Barro–Lee dataset indicates that the US is the female human capital frontier. I use the NLSY79 (1958–1965 cohort) to compute average female education (13.75 years) and reweighted average education (13.92 years) at the frontier. To compute "average human capital", I transform years of education to Hall and Jones's human capital measure and then average across individuals. To compute "aggregate human capital", I average years of education and then perform the Hall and Jones transformation.
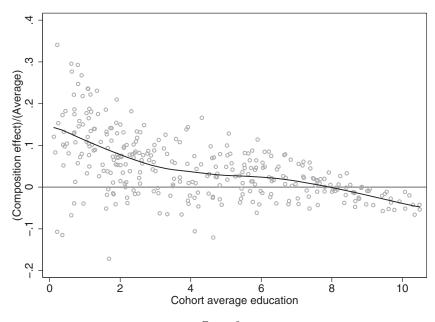
FIGURE 8

Composition effects as a share of average education.

Notes: The curve is a local linear regression with a bandwidth of 1 year. The composition effect is the difference between average education and the counterfactual that would arise if all families had the same number of siblings, with no change to their education. Sample includes all women over the age of 20 years, except for the 2001 Bangladesh survey and the 1996 Nepal survey, which include ever-married women over the age of 30 years. Data source: DHS Sibling Histories

a counterfactual in which both the DHS and the US data are reweighted to remove the composition effect of differential fertility. This effect is expressed proportionally:

$$\text{Effect on distance to the frontier} = \frac{\bar{h}_{US} - \bar{h}_c}{\left(\bar{h}_{US} - \Delta_{comp,US}\right) - \left(\bar{h}_c - \Delta_{comp,c}\right)} - 1, \qquad (6.2)$$

where $\bar{h}_c$ is average human capital (or educational investment) in country c and $\Delta_{comp,c}$ is the composition effect of differential fertility in country c. Continuing in the mold of Figures 7 and 8, column (3) shows results for average years of education. Most estimates are negative, implying that differential fertility *reduces* the skill gap between developing countries and the US. In the most extreme case, the gap in average education between the Republic of Congo and the US is 8.5% smaller than it would have been in the absence of differential fertility within each country. Averaging across all 43 countries, differential fertility reduces the distance to the frontier by 3%. Again, these patterns suggest that differential fertility helped rather than hurt developing countries in the second half of the twentieth century.

Measures of human capital lead to similar conclusions. For comparison with the development accounting literature, columns (4) and (5) use the human capital production function of Hall and Jones (1999): $h = \exp[f(s)]$, where $f(s)$ is a piece-wise linear function of years of schooling s.[34] Column (4) transforms years of schooling to human capital and then averages across women (average human capital), whereas column (5) averages years of schooling across

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34. Specifically, the coefficient on s is 0.134 for $s \in [0,4]$, 0.101 for $s \in (4,8]$, and 0.068 for $s > 8$.

women and then performs the transformation (aggregate human capital). Because the Hall and Jones production function exhibits decreasing returns (and sample countries have low levels of education), either approach leads to effects that are slightly larger in magnitude than the estimates in column (3). On average, differential fertility reduces distance to the frontier in mean human capital by 4%.

7. CONCLUSION

Efforts to understand whether and how distributional considerations play a role in the escape from the Malthusian trap have been stymied by fragmentary evidence on how cross-sectional patterns of fertility and child investment change over the demographic transition. With the goal of filling that gap, this article studies the evolution of these patterns over half a century of birth cohorts in 48 developing countries. The results suggest that the relationships linking income or skill with fertility are initially hump-shaped, with most of the population in the domain with a positive slope. As the economy develops, the peak of the hump shifts to the left, and the skill distribution shifts to the right, such that the associations of income or skill with fertility flip from positive to negative. Mirroring this reversal, children from larger families initially obtain more human capital, but this association flips with economic development. Increases in the aggregate education levels of the parents' generation are by far the most important predictor of the reversal; the data show little role for child mortality rates, GDP *per capita*, sectoral composition, urbanization, and women's labour force participation. Given the unique role of rising aggregate education and the shift of the peak of the fertility–durable goods relationship, the data are most consistent with a theory in which a rising return to child investment leads families further and further down the income distribution to invest.

Because the reversal has gone largely unrecognized in much of the literature on the aggregate effects of differential fertility, that literature has missed an important aspect of the interaction between demography and economic growth. In the mid-twentieth century, fertility differentials by parental income and skill increased average education in most of the countries under study. In the least educated countries, the positive effect of differential fertility amounted to roughly 15% of average education. As aggregate education rose, fertility differentials eventually flipped in many countries, and so too did the effect of differential fertility on average education. Whether the composition effects identified in this article are large enough to play an important role in endogenous growth is a fruitful direction for future research, as are their implications for the evolution of income inequality.

Another fruitful direction for future research concerns policy implications. The results imply that increased inequality may promote economic growth early in the development process but depress it later. But redistributive taxation may not promote growth in either phase. In the Malthusian regime, such a policy may decrease the already small share of children who receive any schooling, while in the modern regime, it may exacerbate fertility differentials, a point emphasized by Knowles (1999). A second policy tool, direct caps on the number of children per women, may have detrimental compositional consequences in the Malthusian regime and beneficial compositional consequences in the modern regime. But if the fertility limit is set sufficiently low, then increases in child investment from re-optimization may swamp these composition effects. Finally, policies that reduce the cost of schooling may increase child investment, but in the Malthusian regime, that increase may be partially offset by a reduction in the population share of the most educated children.

The article's results also relate to a recent literature suggesting a further reversal of differential fertility at higher levels of development. In cross-country data, Myrskylä *et al*. (2009) find that the long-standing negative association between the Human Development Index and the total fertility

rate has turned positive among the most developed countries, although Harttgen and Vollmer (2014) question the robustness of this result. And within the US, Hazan and Zoabi (2015) estimate an emerging U-shape in the relationship between women's education and fertility, such that women with advanced degrees have more children than women with just undergraduate degrees. Taken together, the evidence suggests that the equilibrium relationship between economic resources and fertility may be "N-shaped", with peaks and troughs that shift with the fundamentals of the economy. While these findings on advanced economies reflect mechanisms different from those driving the results of this article, they reinforce the conclusion that the association between economic resources and fertility is not always negative.

APPENDIX TABLE 1

DHS countries and survey years

Fertility Histories	Sibling Histories
Burkina Faso 1992–3, 2010; Burundi 1987, 2010; Cameroon 1991, 2011; Colombia 1986, 1990. 2009–10; Dominican Republic 1986, 1991, 2007; Ghana 1988, 1994, 2008; Haiti 1994–5, 2005–6; India 1992–3, 2005–6; Indonesia 1994, 2007; Kenya 1988–9, 2008–9; Madagascar 1992, 2008–9; Malawi 1987, 2006; Nigeria 1990, 2008; Niger 1992, 2006; Namibia 1992, 2006–7; Peru 1986, 1992, 2006–8; Senegal 1986, 1993, 2010–11; Tanzania 1991–2, 2010; Zambia 1992, 2007; Zimbabwe 1988, 1994, 2010–11.	Afghanistan 2010; Bangladesh 2001; Benin 1996, 2006; Bolivia 1994, 2003, 2008; Burkina Faso 1999, 2010; Burundi 2010; Cambodia 2000, 2005, 2010; Cameroon 1998, 2004; Central African Republic 1995; Chad 1996, 2004; Congo, Dem. Republic 2007; Congo, Republic 2005; Cote d'Ivore 1994; Dominican Republic 2002. 2007; Ethiopia 2000, 2005, 2010; Gabon 2000; Ghana 2007; Guinea 1992, 2005; Haiti 2000, 2005–6; Indonesia 2012; Kenya 1998, 2003, 2008–9; Lesotho 2004, 2009; Madagascar 1992, 1997, 2004, 2008–9; Malawi 1992, 2000, 2004, 2010; Mali 1995, 2001, 2006; Morocco 1992, 2003; Mozambique 1997, 2003; Namibia 1992, 2000; Nepal 1996, 2006; Nigeria 2008; Peru 1992, 1996, 2000, 2003–8; Philippines 1993, 1998; Rwanda 2000, 2005, 2010; São Tomé & Príncipe 2008; Senegal 1992, 2005; Sierra Leone 2008; South Africa 1998; Sudan 2010; Swaziland 2007; Tanzania 1996, 2004, 2010; Togo 1998; Zambia 1996, 2001, 2007; Zimbabwe 1994, 1999, 2005, 2010–11

Supplementary Data

Supplementary data are available at *Review of Economic Studies* online.

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